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An Application of the LSE Methodology in Systems of Nonstationary Variables

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The Wage-Price spiral in Greece: an application of the LSE methodology in systems of nonstationary variables

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Abstract

Wage and price determination in Greece is investigated empirically in the context of a closed system during the period 1975 - 1990. The analysis adopts the "general to specific" methodology, in which the time dependence properties of the series play an important role. After a univariate analysis of the data series with emphasis given to their seasonal behaviour, a multivariate cointegration technique leads to the identification of one long run relationship among the series analysed: a real wage - productivity relationship, with positive unemployment effects. A theoretically reasonable simultaneous equation model (SEM) is finally established by testing for congruence and encompassing against a congruent vector autoregression (VAR) which incorporates the relevant long-run information and is shown to provide better forecasts than a VAR in differences model which constitutes a strong rival model.

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1 Introduction.

The aim of the present paper is to investigate empirically the wage and price spiral in Greece over the post-1974 period, which is characterised by high inflation rates. The modelling strategy adopted in the work follows recent developments in the econometric literature, in the spirit of the "LSE methodology".

The analysis is performed in the context of a closed system which includes hourly wages, consumer prices, hourly productivity and unemployment. Particular attention is paid on the time dependence properties of the series. In a first step, univariate data series analysis is done: it includes firstly examination of the seasonal pattern of the series by applying an ARIMA model-based adjustment technique and by testing for seasonal integration of the series and secondly, testing for integration at zero frequency. In a second step, cointegration at zero frequency is tested in a multivarate framework by using the Johansen (1988) maximum likelihood cointegration technique; the long-run relationships between the variables are identified with emphasis given in the parameter constancy of the relations and the exogeneity/ endogeneity status of the included variables with respect to the long-run parameters of interest.

Then, the "general to specific" methodology is applied in order to select a final simultaneous equation model (SEM) describing the dynamics of the system, while incorporating the long-run information. Its strength is evaluated by its ability to encompass a vector autoregressive representation (VAR) which includes the long-run information and can be considered a congruent representation of the joint distribution of the series of interest. It is also compared with a VAR in differences (DVAR), with special emphasis given to their forecasting ability.

The rest of the paper is organised as follows: The next section describes briefly the econometric methodology adopted for the analysis. Section 3 highlights some characteristics of the Greek labour market institutions and their implications for modelling the wage-price spiral; it also gives descriptive information on the data and defines the sample period for the study. Section 4 presents the univariate time dependence analysis of the series and gives the arguments for the choice of the variables to be used in the modelling. In Section 5 the multivariate cointegration analysis is performed: the long-run relationships are identified and the exogeneity status of the variables is assessed. Section 6 derives the SEM and evaluates its adequacy, while the last section summarises and concludes.

2 The econometric methodology.

The general framework of the "LSE methodology" applied in the present paper is extensively discussed in Hendry (1995), Hendry and Mizon (1990), (1993), Hendry and Richard (1983), Mizon (1995b) and Spanos (1986), (1990a), *inter alia.* Basic concept of the approach in hand is that theories are treated as providing approximations to the observable phenomena without being exact copies of reality. In this context, observed data constitute a sample taken from an on-going real data generation process (DGP) with all its variability and "irrelevant" to the theory features, together with observational errors, while a theoretical model is simply a mathematical formulation of a theory based on simplifying assumptions.

The first step in this modelling strategy is the specification of a statistical model approximating the actual DGP which can be considered to be represented by the joint distribution of the observed variables. The estimated statistical model, has to be shown to be statistically adequate, in the sense that the assumptions defining it are valid. It can then be used as a valid basis against which alternative simplifications can be evaluated[1]. Following this procedure, a final econometric model can be chosen, which, however, has to be shown to be congruent with the available information and to encompass the unrestricted system. Where congruence entails data coherency, constant parameters, valid weak exogeneity of any unmodelled variables for the parameters of interest, consi-

[1]The need for a common statistical framework in the context of non-nested models and encompassing was emphasised by *inter alia* Mizon (1984), (1995a), Mizon and Richard (1986).

stency with *a priori* theory, and data admissibility (Hendry and Richard (1983), Hendry and Mizon (1993)).

In a systems context, Hendry and Mizon (1993), Clements and Mizon (1991) propose the use of a congruent unrestricted vector autoregressive representation (UVAR) specified in levels as a valid representation of the actual DGP allowing for non-stationarities in the variables analysed[2].

In a VAR framework the number of the possible long-run cointegrating relationships between the variables can be defined following the procedure suggested by Johansen (1988), Johansen and Juselius (1990). Then, the identified long-run information can be included in a reparameterisation of the original UVAR, which can be used as a benchmark within which alternative SEMs can be compared. This procedure is discussed in more details below:

The familiar vector equilibrium correction model (VECM) form[3] of a VAR is:

$$\Delta x_t = -\sum_{i=1}^{p-1} \Pi_i \Delta x_{t-i} + \Pi x_{t-p} + \psi D_t + \nu \tag{1}$$

where $\nu_t \sim IN(0, \Sigma)$, and x_t is an $N \times 1$ vector of the time series variables of interest and D_t contains deterministic components (constant, trend, seasonal dummies and event dummies). The order of the VAR p is assumed finite, so that moving average components are excluded and the parameters Π_i, Π, ψ and Σ are assumed constant. Π is the matrix of the long-run responses and if there exist r cointegrating relationships between the variables, is of reduced rank $r < N$. In this case, Π can be expressed as the product of two $N \times r$ matrices α and β'

$$\Pi = \alpha\beta' \tag{2}$$

[2]Monfort and Rabemanjara (1990), propose a similar methodology for stationary series.

[3]The Clements and Hendry (1995) terminology is adopted here, based on the observation that in such reparameterisations the long-run information terms known as "error corrections" first introduced by Davidson *et al* (1978) may play the opposite role when the equilibrium changes, so they should be called "equilibrium corrections".

where β contains the r cointegrating vectors and α is the loadings or adjustment parameters matrix, which contains the loadings with which the cointegrating relationships enter the equations modelling $\Delta \mathrm{x}_t$.

Johansen and Juselius (1990) provide the test statistics to define the rank of the matrix Π and show that testing for linear restrictions on either the parameters of the cointegrating vectors or their loadings, is allowed given that the matrices α and β' are not unique. Therefore, specific meaningful economic restrictions concerning the long-run parameters of interest β_i, as well as restrictions on α (some zero restrictions on α_i correspond to weak exogeneity of the variables for the long-run parameters) can be tested and not imposed *a priori* (for a detailed analysis of exogeneity testing, see *inter alia* Ericsson and Irons(1994)). As shown, standard asymptotically χ^2 likelihood ratio statistics can be used for these restrictions.

The chosen restricted cointegrating vectors can then play the role of equilibrium correction terms in a reduced rank parameterisation of (1), in which, though, the estimated short-run dynamics and the coefficients of the deterministic variables are changed. It is of the form:

$$\Delta x_t = \sum_{i=1}^{p-1} \Pi_i' \Delta x_{t-i} + \phi ECM_{t-1} + \psi' D_t + \nu \tag{3}$$

(3) is a I(0) parameterisation which includes the long-run information of the series behaviour; it has fewer parameters than the original VAR and so it can be referred to as a parsimonious VAR (PVAR). It can then be used as a benchmark within which alternative SEMs can be compared, the advantage of doing so being the use of models who are robust to changes in the sample information (see Mizon (1995b), Clements and Mizon (1991)). The strategy results in a full system of equations, rather than a single reduced form; it thus allows for using the more powerful test of forecasting ability in which predicted values of all variables are used rather than actual values of key variables.

3 The variables set, Greek labour market institutions.

3.1 The variables set.

Empirical work on wage determination has been greatly influenced by the seminal work of Sargan (1964), where he provides one of the earliest forms of an error correction model. Recent work entails application of the notion of cointegration for the estimation and testing of wage long-run equilibrium relationships. (Among others, Hall (1989), Alogoskoufis and Smith (1991), Clements and Mizon (1991), Mizon (1995b), test for cointegrating relationships among labour market variables for the case of the United Kingdom; Kouretas (1993) for the EFTA economies; Juselius (1992), Nymoen (1992) and Psaradakis (1991), for the Danish, the Finnish and the Greek economies respectively).

In the present analysis the set of variables modelled was chosen to be similar to that used by Hall (1989) and Clements and Mizon (1991). The data set covers the unemployment series (U), the hourly earnings in manufacturing (W), the consumer prices (P) and the hourly productivity in manufacturing (HPROD) derived as the rate of the hourly industrial production in manufacturing (defined as the rate Y/H), over the employment (E) in that sector. All data series are quarterly and not seasonally adjusted. Detailed definitions of the series as well as data sources can be found in the Appendix. The natural logs of the series are used (this is done, for the unemployment rate variable case, in order to have the models' fitted and predicted values bounded between 0 and 1). Throughout the paper, lower-case letters which refer to the variables signify logarithms of capitals, and D denotes the first difference operator.

The choice of variables is rather restricted: the purpose of the paper is to model the wage and price spiral focusing on the labour market sector. Therefore, other possible determinants of the price inflation (such as exchange rates, import prices, or the public deficit) are not included in the analysis. This is also done in order to keep the system manageable, given that inclusion of many variables would mean too few degrees of freedom for statistical inference, unless if this was done at the expense of introducing a range of exogeneity assumptions. The analysis is here constrained to be done on the context of a closed system involving the variables mentioned above.

3.2 Changes in regime, labour market institutions.

The plots of the series used are given in Figures 1 and 2. The series presented in Figure 1 are the quarterly consumer price inflation Dp, the annual consumer price inflation $D4p$, the quarterly real wage inflation Drw, the annual real wage inflation $D4rw$, the real wage rw, and the unemployment series U. The wage variables refer to the manufacturing sector, and the graphs of a number of variables characterising the performance of that sector, are given in Figure 2[4]. These variables are: production Y, employment E, weekly hours worked H, productivity $PROD$ and hourly productivity $HPROD$.

In the present work we try to model the wage and price spiral for the post-1974 period given that it is characterised by high inflation rates as also shown in the Dp, $D4p$ graphs. In particular, while lower than the OECD average for the fifteen years before 1974, inflation rate rose sharply after the first oil price shock and has remained in the highest positions in the OECD and EEC area from then on. 1974 is also the year when the military regime fell, resulting to a number of changes in the labour market institutions.

The analysis is extended up to 1990.2. This is done so, because at this quarter, a number of restrictive policies taken by the newly elected conservative government (in an effort for the Greek macroeconomic variables to converge to the Maastricht standards), resulted in radical but overestimated changes in the behaviour of basic variables characterising the manufacturing sector performance. More analytically, in 1990.2 the

[4]Even though the manufacturing sector share of GDP is quite low, it is considered to be indicative of the labour market developments and important for the wage formation in the years under consideration.

automatic wage indexation scheme was abolished, policies for subsidysing loss making companies were stopped and privatisations were started. As shown in the graphs of the relevant series in Figure 2, the measures result to a considerable fall in the employment in manufacturing which is not, though, accompanied by the same fall in production: this results to an impressive rise in labour productivity. However, this productivity swing is overestimated, given that it reflects the closure or restructuring of a number of low productivity "problematic" enterprises. In addition, in the 90's manufacturing accounts for only the 17% of GDP share compared to the 25% at the 70's, while emphasis is now given by the authorities to the development of the touristic service sector[5].

Analysis of the wage and price inflation determination during the 1975 - 1990 period is of interest given that it covers two different political regimes with different weights placed on inflation control: the conservative government episode until 1981.3 is succeeded by an episode during which the socialist party is in power, even though this second episode is characterised by a switch towards more restrictive policies in 1985.4, when a stabilisation program is put into practice. 1975 - 1990 also covers a number of events and institutional changes that affected the performance of the Greek econony. From March 1975 the drachma is not linked to the dollar but a managed exchange rate regime is followed. In 1979 the econony has to deal with the second oil price shock. In January 1981 Greece becomes an EEC member country. In January 1982, the socialist government introduces a formal (but not full) automatic wage indexation scheme, which remains in practice until 1990, excluding the 1986 - 1987 stabilization program period. Finally, in October 1985, a stabilisation program (including devaluation of the drachma by 15% in order to raise Greek competitiveness) is put into effect.

[5]Attempts to model the whole 1975 - 1995 period resulted in models and cointegrating relationships with non-constant parameters. An alternative would be modelling the price-wage spiral in the whole period using a reduced system including wages, prices and unemployment (see Sideris (1996)).

3.3 Descriptive analysis.

The effects of these shifts are evident in the graphs of the series. The inflation rate rises considerably in 1980 and remains high for the following three years as a result of the expansionary policies that were followed during this period. First, by the conservative government in order to accomodate the second oil price shock in 1979 and in the pre-election 1981 year and then by the socialist government elected in October 1981 for the first two years they were in power. It reaches its highest observed point in 1986.1 because of the drachma devaluation in the previous quarter, while relatively low rates are observed during the stabilisation program period 1986 - 1987.

Hourly real wages show an upward trend for the period until 1985.4, decrease during the 1986 - 1987 stabilisation scheme period, and remain relatively stable from then on, indicating the shift to more restrictive policies. The unemployment rate is characterised by an upward trend especially for the period until 1986; the seasonal pattern is also very strong and this can be attributed to the fact that employment in Greece is highly related to the developed touristic service sector.

Production, employment and productivity in manufacturing show an upward trend up to 1981.1 and remain relatively constant for the rest of the period, despite the on average expansionary policies of the 80's; the evidence makes more apparent the structural inefficiencies of the productive sector which could not respond to demand increases. This probably indicates that this sector which operated in an environment of protectionism for years could not adjust promptly in the EC competitive environment. Average weekly hours worked show a downward trend from the beginning of the examined period till the mid 80's when they reach quite low levels; they follow a stable path after 1988, reflecting a stabilisation in the labour market conditions. Reflecting the pattern of the above series, hourly productivity remains also relatively stable during the 80's.

As described above, the behaviour of the series of interest has been strongly influenced by particular policy shift events that took place in certain time points and thus may support inclusion of dummy variables. As advocated by Clements and Mizon (1991), inclusion of dummies is preferable to the enlargement of the number of explanatory variables, given that the sample size is relatively small. This should be kept in mind while extending the analysis in Section 5, where a UVAR is formed.

In Figure 3, the graphs of more than one series of interest adjusted for mean and variance are given together, in an attempt to investigate visually possible relationships among them. A seasonally adjusted series for unemployment, su, (obtained by an ARIMA- model-based techique as described in Section 4) is used instead of the seasonally unadjusted one in order to make the changes of the pattern of unemployment more evident. The first phenomenon that could be naively observed in graph (a) which depicts the annual inflation-unemployment relation is a Phillips curve relationship for the years 1980 - 1985. This, however, would be wrong: the modest decrease in inflation and the rise in unemployment are not the results of restrictive policies (which were quite expansionary during this period). In addition, as it is also shown by graph (b) this period is characterised by simultaneous increase in real wages and unemployment. The evidence can probably be explained by the insiders - outsiders theories that claim stronger interest of the powerful insiders (who care for the welfare of their employed members) for increase in real wages than decrease in unemployment, and the assumption of real wage rigidity. The argument is strengthened by graph (c) where employment in manufacturing remains stable, despite the increase in unemployment. This indicates mainly changes in the structure of unemployment but it also reflects the fact that the manufacturing sector could not absorbe new entrances in the labour force; the explanation being twofold: i) it could not respond to positive demand shocks as functioning in the new competitive EC environment, and ii) the rise in real wages did not allow for new working places. Finally, graph (d) shows that real hourly wage was increased on top of the increase in hourly productivity in the periods 1975 - 1977 and 1982 - 1985.

Summarising, we argue that analysis of the wage-price spiral during the 1975 - 1990 period is of interest, given that the period is characterised by high inflation rates and covers different policy regimes, while, as also evidenced by the graphs of the series, an important policy change takes place in 1985.4 signaling a switch to more restrictive policies.

4 Univariate analysis of the time properties of the series.

4.1 Characterisation of the seasonal pattern.

The univariate analysis of the series entails initially thorough investigation of their seasonal pattern. Issues such as the significance of the seasonal component on the evolution of a series, whether seasonality follows a constant pattern or not, or to what extend the series are integrated at seasonal frequencies are of importance for the modelling of the closed system. As shown by the graphs of the series, the presence of seasonality is evident for at least the series of productivity and unemployment. In the present application, we therefore first use an ARIMA-model-based method in order to estimate the seasonal component of the series and then test for integration of the series at a seasonal frequency.

4.1.1 Estimation of the seasonal component.

Estimation of the seasonal components of the series is performed by applying the SEATS (Signal Extraction in ARIMA Time Series) (Maravall and Gomez (1994a)) programme. Given that it assumes a linear time series model with Gaussian innovations, it was used in companion with TRAMO (Time series Regression with ARIMA noise, Missing observations and Outliers) (Maravall and Gomez (1994b)) which played the role of a preadjustment program[6].

[6]TRAMO was used in order to identify and correct for outliers in the series; it actually detected outliers at 1985.4 and 1986.1 for the price series. The corrected series was therefore used for the seasonal adjustment of the price variable; for the rest of the cases, the original raw series are used for estimation of the seasonal components.

Table 1: **Estimation of the seasonal component with SEATS.**

			AIRLINE model				
	coeff.		Missp.	Tests			
Variable	δ_1	δ_4	st. err.[1]	DW	skewn.	kurt.	$N(\chi^2(2))$
p	0.219	-0.257	0.905	1.878	0.311	2.656	1.475
w	0.151	-0.746	0.211	1.999	0.241	3.255	0.916
hprod	-0.604	-0.613	0.288	1.903	-0.054	3.741	1.709
u	0.243	-0.240	0.104	2.086	-0.306	3.848	3.831

[1] Standard error in 10^{-1}

SEATS is an ARIMA-model-based method for decomposing a series into its unobserved time components: trend, seasonal, cyclical and irregular components and it is used for seasonal adjustment of economic time series. The programme is fitting by default the so-called airline model (see Box and Jenkins (1970)) which provides a decent fit to the series according to Gomez and Maravall (1994). The airline model is given by:

$$(1-L)(1-L^4)X_T = (1-\delta_1 L)(1-\delta_4 L^4)\epsilon_t + \mu \tag{4}$$

where ϵ is a white noise innovation and μ is a constant. SEATS uses a model-based technique and therefore provides with diagnostics that allow for evaluation for the fit of the model. It also provides an estimate of the seasonal pattern, and the weights by which it contributes to the estimate of the series. In Table 1, the diagnostics of the fitted models are reported, together with estimates of the parameters δ_1, which is related to the stability of the trend component, and δ_4 which is related to the stability of the seasonal component.

However, analytical results for the TRAMO preadjustment procedure are not given for space reasons.

In Figure 4, the estimated trends, seasonal components and seasonally adjusted series are presented, while Figure 5 presents the weights by which the seasonal pattern is contributed to the estimated series. The seasonal component is quite unstable for the cases of the unemployment and price series. Finally, as shown in Figure 5, the seasonal pattern plays a very important role for the evolution of the unemployment series, while it has minor impacts for the evolution of the rest of the series. The evidence advocates the use of a seasonally adjusted series (as suggested *inter alia* by Hendry (1995), p. 559-565) instead of the raw series for the unemployment variable when going on with the modelling of the system. Nevertheless, further investigation of the seasonal pattern of the whole group of series by testing for seasonal integration is attempted before proceeding with the multivariate cointegration analysis.

4.1.2 Testing for seasonal integration.

The stochastic process X_t is integrated of order (n, s), or $I(n, s)$, if the series is stationary after first period differencing n times and seasonal differencing s times. The most used test for seasonal integration is the Hylleberg, Engle, Granger and Yoo (1990) (HEGY) test, which considers all the possible seasonal roots of the generating process. It essentially allows the null hypothesis of $I(1,1)$ to be tested against the alternatives of $I(0,1)$ and $I(0,0)$ by making use of the following regression equation:

$$A(L)Y_{4t} = \gamma_1 Y_{1t-1} + \gamma_2 Y_{2t-1} + \gamma_3 Y_{3t-2} + \gamma_4 Y_{3t-1} + \epsilon_t \tag{5}$$

where Y_{it} are transformations of the time series X_t:

$$Y_{1t} = (1 + L + L^2 + L^3)X_t \tag{6}$$

$$Y_{2t} = -(1 - L + L^2 - L^3)X_t \tag{7}$$

$$Y_{3t} = -(1 - L^2)X_t \tag{8}$$

Table 2: **HEGY tests.**

Variable	Sample	Det	γ_1	γ_2	γ_3	γ_4
p	76.1-90.2	I, SD	-0.566	-4.131*	-3.209	-2.404*
		I, Tr, SD	-2.649	-3.913*	-3.127	-2.011*
w	76.1-90.2	I, SD	-1.869	-3.726*	-4.031*	-1.579
		I, Tr, SD	-0.721	-3.653*	-4.021*	-1.583
hprod	76.1-90.2	I, SD	-1.983	-3.101*	-1.034	-4.455*
		I, Tr, SD	-2.477	-3.321*	-0.857	-4.283*
u	76.1-90.2	I, Tr	-3.061	-1.308	-0.563	-1.018
		I, Tr, SD	-2.876	-1.640	-1.067	-0.403
su	76.1-90.2	I, SD	-0.273	-3.190*	-4.988*	-2.919*
		I, Tr, SD	-2.712	-3.397*	-4.137*	-3.408*

$$Y_{4t} = (1 - L^4)X_t \tag{9}$$

The order of the $A(L)$ polynomial is obtained through augmenting the basic regression parsimoniously by lags of Y_{4t} to ascertain an *iid* error process ϵ_t. Deterministic terms such as an intercept (I), an intercept and a trend (I, Tr), an intercept and seasonal dummies (I, SD), or an intercept, a trend and seasonal dummies (I, Tr, SD) can be added to the regression.

Stationarity of X_t requires that γ_1, γ_2, and either γ_3 or γ_4 are non-zero. If $\gamma_1 = 0$, whilst γ_2, and either γ_3, or γ_4 are non-zero, $I(1,0)$ behaviour is implied. If $\gamma_2 = 0$, X_t has a unit root at the biannual frequency, whilst $\gamma_3 = 0$, and/or $\gamma_4 = 0$, imply a unit root at an annual frequency. This last hypothesis can be tested by either an F test for $\gamma_3 = \gamma_4 = 0$, or a two-sided $t - test$ for $\gamma_4 = 0$, followed by a one-sided $t - test$ for $\gamma_3 = 0$, if $\gamma_4 = 0$, is not rejected. The finite sample distributions of the test statistics testing the above hypotheses are tabulated in Hylleberg et al (1990).

The results of the HEGY test are reported in Table 2. The I(1) property at zero frequency for every series is stated by the t_{γ_1}, statistic.

Then, for the series p, w, and $hprod$, the assumptions $\gamma_2 = 0$, and either $\gamma_3 = 0$, or $\gamma_4 = 0$, are rejected at a 5% level, implying a $I(1,0)$ behaviour: the series are not seasonally integrated. However, for the case of the unemployment series, the presence of seasonal unit roots cannot be rejected; in particular, the HEGY tests strongly indicate unit roots at both the annual and the biannual frequency. (Given that no other series turns out to be an integrated seasonal process, there is no ground for testing for seasonal cointegration).

Finally, as Hylleberg *et al* (1990) suggest, it would make sense to use a filtered series in place of the seasonally integrated u, when testing further for cointegration at zero frequency with the rest of the series. Therefore, the seasonally adjusted series su estimated by the SEATS technique as described in the previous subsection, is going to be used for the multivariate analysis[7].

The graph of su is already shown in Figure 3. As expected, the HEGY tests performed for the su series which are reported at the low part of the Table 2 do not indicate the presence of seasonal unit roots.

4.2 Testing for integration at zero frequency

The by now well known univariate augmented Dickey and Fuller (1981) (ADF) tests are applied to check for the presence of unit roots at zero frequency. These tests rely on the rejection of the hypothesis that a process is driven by a random walk against the alternative of stationarity. The results are reported in Table 3[8]. The regressions include a constant and seasonals, when testing for the unadjusted series to account for any deterministic seasonality. Using a 1 % significance level, the data clearly

[7]Note that Ericsson, Hendry and Tran (1993) suggest that use of either adjusted or unadjusted series leads to similar results in terms of the estimated cointegrating vectors; in the present paper, though, it was decided not to use the non adjusted unemployment series given that it was found to entail seasonal unit roots and to contain a very strong and unstable seasonal pattern.

[8]The results reported are obtained using the PC-GIVE module of the PC-GIVE version 8.1, system of computer programs (see Doornik and Hendry (1994)).

Table 3: **Augmented Dickey-Fuller Tests.**

Variables	t(ADF)	lag length
p	-0.16104	3
w	-1.8068	3
hprod	-2.2039	4
su	-0.91873	4
u	-0.35346	4
Dp	-2.9271*	3
Dw	-4.6036**	3
Dhprod	-5.2291**	4
Dsu	-3.5718**	4
Du	-3.9942**	4

* significant at 5% level

** significant at 1% level

reject the first order integration hypothesis in favour of a stochastically stationary alternative in the case of *Dhprod*, *Dsu* and *Dw*, while for the levels of all four variables, show no evidence against the I(1) representation. The presence of unit root is, however, marginally rejected for the case of the *Dp* series, giving evidence that it may be integrated of order 2.

However, the D-F unit root tests are low power tests; in particular, their power is likely to be very low for values of the autoregressive parameter less than, but close to unity. In addition, unit roots are not invariant to changes in the information set relative to which they are defined. (see Spanos (1990b)). Hence, a multivariate analysis of the time dependence properties of the series seems to be more appropriate.

5 Multivariate cointegration analysis.

5.1 The unrestricted VAR.

An unrestricted fifth order autoregressive system (UVAR) for the vector $x'_t = (p, w, hprod, su)$ containing also a constant term and centred seasonal dummies, was initially estimated for the period 1975.1 to 1990.2 using multivariate least squares. Lower order UVAR systems were evaluated against it by using likelihood ratio tests, provided there were no autocorrelated residuals in the specifications. A fourth-lag system was finally found to adequately capture the dynamics.

However, there remained evidence (shown by Chow tests for parameter constancy) of two substantial but explainable outliers: in 1975.2 for the switch to a managed exchange rate regime and in 1985.4 for the change in the economic policy which included a drachma devaluation. The effect of the two outliers was eliminated by the use of the dummies D75.2 which takes the value 1 in 1975.2 and 0 elsewhere and D85.4 which takes the value 1 in 1985.4 and 0 elsewhere. Since the two dummy variables should not have a long-run effect on any of the modelled variables, they are entered unrestricted in the VAR. They both turned out to be significant at a 5% level with obtained t-values 2.8323 (0.0381*) for D85.4, and 4.4626 (0.0045**) for D75.2; in addition, their absence would mean nonnormality for the residuals[9].

The descriptive statistics of the unconstrained fourth order VAR system are presented in Table 4. First, single equation diagnostics are reported: the *AR* Lagrange multiplier (LM) statistic for residual serial independence across five lags of the autocorrelation function , the *ARCH* LM test statistic testing the null of no autoregressive conditional hete-

[9] A number of other impulse dummies to account for events that have possibly influenced the behaviour of the series (for the periods 1979.1, 1981.1 and 1983.1 to account for the oil price shock, Greece becoming an EEC member and a first drachma devaluation), were also included for the specification of the system, but they turned out to be statistically insignificant and therefore they were not kept in the final specification of the unrestricted VAR.

roscedasticity and the N statistic testing the null of normal skewness and kurtosis. Second, test statistics for vector autoregressive residuals $vecAR$ and vector normality $vecN$ (see Doornik and Hendry (1994), for definition of these test statistics). There is no evidence for misspecification of the residuals of the estimated VAR.

Furthermore, the parameter constancy assumption was assesed by the sequence of forecast Chow tests against the end point of the sample (not shown here in order to save space): the tests imply that the parameters remain constant over the examined period. Similar evidence is also borne out by the tests for predictive failure F_1, F_2, F_3 for the last eight observations (for details for the tests see Doornik and Hendry (1994)): according to these tests, the estimated parameters remain reasonably constant over the period 1988.3 - 1990.2 (it is only F_1 (32,33) that rejects parameter constancy). Finally, inspection of the residual correlations suggests that there is a modest correlation between $hprod$ and w and $hprod$ and su but the correlations between the residuals of the rest of the equations are negligible.

5.2 Cointegration analysis.

5.2.1 Identification of the cointegration space rank.

Having established a VAR system which provides an adequate characterisation of the data structure, and fulfills the required assumptions (residuals which are serially uncorrelated, homoscedastic, and normally distributed and (relatively) constant parameters), we can go on by examining the time dependence of the data series within a multivariate framework. The Johansen maximum likelihood technique in which the order of cointegration of the system is examined conditional upon the short-run dynamics of the Δx_t process and the seasonal dummies, is therefore applied.

The estimated eigenvalues and the results of the two rank tests, are given in the upper part of Table 5. The largest eigenvalue which is involved in the maximisation of the loglikelihood function with respect to

Table 4: **UVAR Diagnostic Statistics.**

Equation standard deviations

w	p	su	hprod
0.02612	0.01141	0.04075	0.01913

Equation tests

Variable	Statistic	Value	p-value
w :	AR 1- 4F(4, 36) =	0.6392	[0.6379]
p :	AR 1- 4F(4, 36) =	0.6460	[0.6333]
su :	AR 1- 4F(4, 36) =	1.9937	[0.1163]
hprod :	AR 1- 4F(4, 36) =	1.0104	[0.4150]
w :	Normality $\chi^2(2)$=	1.8138	[0.4038]
p :	Normality $\chi^2(2)$=	1.6403	[0.4404]
su :	Normality $\chi^2(2)$=	0.2362	[0.8886]
hprod :	Normality $\chi^2(2)$=	0.0359	[0.9822]
w :	ARCH 4 F(4, 32) =	0.5589	[0.6940]
p :	ARCH 4 F(4, 32) =	1.9461	[0.1268]
su :	ARCH 4 F(4, 32) =	0.9505	[0.4478]
hprod :	ARCH 4 F(4, 32) =	0.4821	[0.7486]

Vector tests

VecAR	1-4 F(64, 84) =	1.3204	[0.1157]
VecN	$\chi^2(8)$=	3.9949	[0.8576]

Parameter constancy forecast tests: sample 1988.3 to 1990.2

F_1 (using Ω)	F(32,33)=	3.1140	[0.0008]**
F_2 (using V[e])	F(32,33)=	1.4782	[0.1346]
F_3 (using V[E])	F(32,33)=	1.5084	[0.1227]

Table 4 (continued):

Correlation of residuals				
	w	p	su	hprod
w	1			
p	0.104	1		
su	-0.095	0.037	1	
hprod	0.195	0.041	-0.161	1

β is quite large (0.33) and turned out to be clearly significantly different from zero (at a 1% level of significance) on the basis of the trace statistic, while its significance is just marginally rejected by the maximum eigenvalue statistic at a 5% level of significance[10]. It was then decided to proceed based on the assumption of one cointegrating vector in the system[11].

The graphs of the cointegrating vectors and the recursive estimated eigenvalues are given in Figure 6. The eigenvalue corresponding to the first cointegrating vector takes a large value and is essentially constant.

5.2.2 Identification of the long-run structure.

The low part of Table 5 records estimates of the standardised eigenvectors and their corresponding loadings of the four variable VAR. An

[10] Critical values of the distributions of the test statistics used are reported in Osterwald-Lenum (1992).

[11] Actually, the second eigenvalue takes a relatively high value and is significant different from zero at a 10% level of significance on the basis of the trace statistic. In addition, as Kostial (1994) indicates, in the case of systems with small eigenvalues of the signal-noise ratio matrix, the Johansen tests tend to underestimate the rank of the cointegrating space in small samples. Therefore, initial analysis was performed based on the assumption of two cointegrating vectors. However, testing for a number of alternative structural restrictions in order to identify two long-run relations among the variables, turned out meaningless: no pair of reasonable economic relationships was accepted by the data set. The analysis was concequently decided to be continued based on the assumption of one cointegrating vector.

Table 5: **Cointegration Results.**

Cointegration analysis 1975 (1) to 1990 (2).

eigenvalue	loglik	rank
	961.569	0
0.333639	974.153	1
0.237604	982.563	2
0.149370	987.578	3
0.030014	988.523	4

H_0:rank=p	Max Eigen.	95%	Trace	95%
$p = 0$	25.17	27.1	53.91**	47.2
$p \leq 1$	16.82	21.0	28.74	29.7
$p \leq 2$	10.03	14.1	11.92	15.4
$p \leq 3$	1.889	3.8	1.889	3.8

Standardised eigenvectors β_i':

w	p	su	hprod
1.000	-0.651	-0.166	-6.951
-2.297	1.000	1.545	8.969
4.109	-4.683	1.000	-10.13
0.123	-0.355	0.172	1.000

Standardized adjustment coefficients α_i:

w	0.059741	-0.023563	-0.048449	0.008061
p	-0.019646	-0.010165	-0.0019498	0.028493
su	-0.080264	-0.107540	0.0023654	-0.033459
hprod	0.067108	-0.018053	0.017363	0.009066

Table 6: **Testing for structural restrictions.**

Hypothesis					$\chi^2(dof)$	p-value
	(w	p	hprod	su)		
H_1:	(1	0	0	0)	10.142 (3)	[0.0174] *
H_2:	(0	1	0	0)	11.565 (3)	[0.0090] **
H_3:	(0	0	1	0)	10.301 (3)	[0.0162] *
H_4:	(0	0	0	1)	12.308 (3)	[0.0064] **
H_5:	(1	-1	a	b)	0.6452 (1)	[0.4218]
H_6:	(1	-1	a	0)	3.5469 (2)	[0.1698]
H_7:	(1	-1	0	b)	3.0202 (2)	[0.2109]
H_8:	(1	-1	-1	b)	3.4631 (2)	[0.1770]
H_9:	(0	0	a	1)	3.5542 (2)	[0.1691]
H_{10}:	(1	-1	-1	0)	4.6116 (3)	[0.2025]
H_{11}:	(1	-1	-1	-0.11)	3.8411 (3)	[0.2791]

examination of the (first) cointegrating vector reported, shows that a direct interpretation is not straightforward. An interesting outcome is that w and p come out with coefficients which are quite close in size to each other and have opposite signs, probably implying a long run relationship between real wage, productivity and unemployment. Nevertheless, further investigation on the identification of the cointegrating vector by testing for possible theoretical assumptions seems to be necessary. A number of theoretical assumptions and their test outcomes are given in Table 6: the likelihood ratio tests reported are asymptotically distributed as χ^2 with the appropriate degrees of freedom given in parentheses.

The first four hypotheses imply stationarity of the individual series. They are all rejected by the given data set, a result which is in line with the univariate testing. The fifth hypothesis H_5 tests for equal in size and opposite in sign w and p coefficients: it implies cointegration between real wage, unemployment and productivity. The relevant likelihood ratio test is assymptotically distributed as $\chi^2(1)$ and H_5 is accepted by the data.

H_6 tests for cointegration between real wage and productivity. H_7 tests for cointegration between real wage and unemployment. H_8 is concerned with the question whether real wage around the productivity trend cointegrates with unemployment, while H_9 implies a long-run relationship between unemployment and productivity. Finally, H_{10} tests for a one to one real wage-productivity relationship. Hypotheses H_6 - H_9 are evaluated by $\chi^2(2)$ tests while H_{10} by a $\chi^2(3)$ test. All hypotheses H_6 - H_{10} are accepted by the data set for p-values which are close to each other. However, theoretical considerations led to the choice of H_8 (which is accepted for the second high p-value among the hypotheses which are assymptotically distibuted as $\chi^2(2)$) as possibly expressing best the underlying relationship.

In addition, when the theoretical assumption is tested in H_{11} assuming that the unemployment coefficient takes the value of -0.11, it is accepted for a p-value of 0.2791, which is the highest p-value obtained for the restrictions implied by the hypotheses H_6 - H_{11}. It was therefore decided to continue the analysis choosing H_{11} as the hypothesis characterising the long-run behaviour of the variables of the given data set. It is of the form:

$$\beta_1 : w_t - p_t - hprod - 0.11su_t \tag{10}$$

It expresses a reasonable positive relationship between real wage and productivity, implying that the wage earners get the share of the productivity growth, with positive unemployment level effects. The positive sign in the unemployment coefficient reflects the fact that the period examined is characterised by quite expansionary policies which included wage increases, but did not result in rises in employment. Such a phenomenon can be due to insiders - outsiders effects, real wage rigidities and inability of the productive sector to react to positive shocks because of labour market rigidities (firing, hiring costs) and the fact that it had to function in the new competitive EEC environment[12].

[12]See Demekas and Kontolemis (1996) for similar arguments in a detailed analysis

Table 7: **Tests for weak exogeneity restrictions.**

Hypothesis	$\chi^2(4)$	p-value
H_{21}: $\alpha_{11} = 0$: w. exogeneity for w:	7.8332	0.0979
H_{22}: $\alpha_{21} = 0$: w. exogeneity for p:	9.6448	0.0469*
H_{23}: $\alpha_{31} = 0$: w. exogeneity for $hprod$:	7.1848	0.1264
H_{24}: $\alpha_{41} = 0$: w. exogeneity for su:	20.074	0.0005**

5.2.3 Tests for weak exogeneity.

Having identified the structure of the cointegrating vector, the analysis can proceed by investigating the exogeneity/endogeneity status of the variables involved. The outcomes of a number of weak exogeneity tests as formed in a multivariate cointegrating framework are reported in Table 7.

Hypotheses H_{21}, H_{22}, H_{23} and H_{24} test respectively for weak exogeneity of wage, price, productivity and unemployment, with respect to the long-run parameters of interest. H_{22} and H_{24} are rejected, implying that prices and unemployment are possibly the endogenous variables in the long-run relationship. The result makes sense, if we take into consideration that during the period, wages were effected to a large extend by trade unions - government negotiations, while productivity is also determined by factors outside the wage determination process.

of unemployment formation and persistence in Greece during the same period.

6 The final model

In the present section, a VAR which models the short-run dynamics including the long-run information (which is known as parsimonious VAR, (PVAR)) is initially estimated. It constitutes the general model within which two nested models can be evaluated: a SEM which simplifies the dynamics of the general formulation and a VAR in differences of the series (DVAR) popular in time series analysis of non stationary series. The two models are compared by considering:

i) their congruency,

ii) their ability to encompass the PVAR,

iii) the constancy of their parameters,

and iv) their forecasting power.

6.1 Encompassing the PVAR

6.1.1 The PVAR

On the basis of the information about the long-run solution to the system, obtained through the cointegration analysis described above, a transformation of the initial system was further decided. The original VAR is transformed into a simplified, yet congruent I(0) representation, by differencing and using the cointegration information. Accordingly, a VAR for the series Dp, Dw, $Dhprod$ and Dsu was estimated, using 4 lags of the series and the cointegrating vector β_1 included as lagged endogenous variable denoted as ecm_{t-1}; in the model no dummies were kept given that they did not turn out to be significant or to improve its diagnostics.

The transformed I(0) system has 4 fewer parameters than the original system and so can be referred to as a parsimonious VAR (PVAR) (see Clements and Mizon (1991), Mizon (1995b)).

Table 8: **PVAR Diagnostic Statistics.**

Equation residual standard deviations

Dw	Dp	Dsu	Dhprod
0.02841	0.01160	0.04058	0.02629

Equation tests

Variable	Statistic	Value	p-value	
Dw :	AR 1- 4F(4, 36) =	0.6215	[0.6501]	
Dp :	AR 1- 4F(4, 36) =	0.6531	[0.6284]	
Dsu :	AR 1- 4F(4, 36) =	3.2881	[0.0214]	*
Dhprod:	AR 1- 4F(4, 36) =	0.5136	[0.7261]	
Dw :	Normality $\chi^2(2)$=	3.6722	[0.1594]	
Dp :	Normality $\chi^2(2)$=	0.6262	[0.7312]	
Dsu:	Normality $\chi^2(2)$=	0.0092	[0.9954]	
Dhprod:	Normality $\chi^2(2)$=	5.1941	[0.0745]	
Dw :	ARCH 4 F(4, 32) =	0.3053	[0.8722]	
Dp :	ARCH 4 F(4, 32) =	2.0878	[0.1055]	
Dsu:	ARCH 4 F(4, 32) =	0.5305	[0.7141]	
Dhprod:	ARCH 4 F(4, 32) =	0.6247	[0.6483]	

Vector tests

VecAR	AR 1-4 F(64, 84) =	1.2926	[0.1345]
VecN	$\chi^2(8)$=	7.8759	[0.4457]

Table 8 (continued):

Correlation of residuals

	Dw	Dp	Dsu	Dhprod
Dw	1			
Dp	0.116	1		
Dsu	0.088	0.178	1	
Dhprod	0.208	0.159	0.080	1

It can be still considered as a congruent parameterisation of the data process as can be seen by the misspecification test outcomes reported in Table 8. The only evidence of noncongruence comes from the autocorrelation statistic for the Dsu equation, which rejects non-autocorrelation but only marginally (p=0.0214); in addition, recursive break-point Chow tests (not shown for economy of space), reveal that the estimated parameters remain reasonably constant over the estimation period. Even though there is scope for simplifying the PVAR specification given that not all the variables included are significantly different from zero, we decided to keep it in this form, so that alternative specifications can be evaluated according to their ability to encompass it.

6.1.2 The DVAR.

The DVAR model corresponds to a model of the form (3) with $\phi = 0$. It is a popular model within the time series analysis tradition (see Box and Jenkins (1970)) and it provides with good forecasts. The diagnostic statistics for the DVAR are presented in Table 9, and indicate that it is well specified.

A LR statistic testing for the overidentifying restrictions implied by the DVAR, which is asymptotically distributed as $\chi^2(4)$ takes the value of 12.128 (p=0.0164*), which rejects the assumption that it parsimoniously encompasses the PVAR. Hence, within sample the PVAR is preferred to the DVAR.

Table 9: **DVAR Diagnostic Statistics.**

Equation residual standard deviations

Dw	Dp	Dsu	DLhprod
0.02811	0.01206	0.04214	0.02624

Equation tests

Variable	Statistic	Value	p-value
Dw :	AR 1- 4F(4, 37) =	0.65957	[0.6240]
Dp :	AR 1- 4F(4, 37) =	0.61438	[0.6550]
Dsu:	AR 1- 4F(4, 37) =	0.95578	[0.4431]
Dhprod :	AR 1- 4F(4, 37) =	0.77913	[0.5459]
Dw :	Normality $\chi^2(2)$=	3.8706	[0.1444]
Dp :	Normality $\chi^2(2)$=	2.3159	[0.3141]
Dsu:	Normality $\chi^2(2)$=	0.53982	[0.7634]
Dhprod :	Normality $\chi^2(2)$=	4.8056	[0.0905]
Dw :	ARCH 4 F(4, 33) =	0.3987	[0.8081]
Dp :	ARCH 4 F(4, 33) =	2.2399	[0.0859]
Dsu:	ARCH 4 F(4, 33) =	0.59138	[0.6712]
Dhprod :	ARCH 4 F(4, 33) =	0.88429	[0.4840]

Vector tests

VecAR	AR 1-4 F(64, 88) =	1.1196	[0.3092]
VecN	$\chi^2(8)$=	10.108	[0.2575]

Table 9 (continued):

Correlation of residuals

	Dw	Dp	Dsu	Dhprod
Dp	0.102	1		
Dsu	0.065	0.211	1	
Dhprod	0.272	0.104	0.031	1

6.1.3 The SEM

Then, alternative simultaneous equations models have been compared by their ability to parsimoniously encompass the PVAR. Among them, the one presented below has been chosen based on simple economic theory considerations, the results of previous relevant studies (for recent works, see Alogoskoufis (1986), (1992), Psaradakis (1991)) and statistical criteria[13]. The model is estimated for the period 1975.2 - 1990.2 using full information maximum likelihood (FIML). It is presented in Table 10.

Wage inflation appears to be influenced mainly by its past values, while price inflation also has a reasonable positive and significant impact on it. Unemployment growth has an overall negative impact on it, implying probably that in the short-run a rise in unemployment has negative effects on nominal wage claims.

The second equation of the SEM shows consumer price inflation to be significantly positively influenced by the history of the process, together with the wage inflation which has a lower but positive and significant impact. The equilibrium correction term has a low but significant effect and it enters with a sign that rules out disequilibrium in the long-run - in line with the interpretation given by Davidson et al (1978).

Unemployment growth is greatly influenced by the history of the process; it is also positively related to wage inflation, result which supports again the long-run cointegrating relationship. It implies that increases in nominal wages would often take place at the expence of decreases in employment, as suggested by micro-based labour market models. Growth in productivity has negative effects, indicating that a rise in hourly productivity works as a motivation for further increase in employment. Finally, the equilibrium correction term has a strong significant positive effect.

Growth in hourly productivity is mainly determined by its past

[13]The "second" powerful simultaneous equations model has similar specification with the one chosen for all but the wage inflation equation. Theoretical considerations, together with the fact that it had lower predictive power, (even though it is wellspecified), led us to choose the one reported.

Table 10: **Simultaneous equation model FIML estimates.**

Variable	Coefficient	t-value	t-prob
Equation for Dw			
Dw_{t-1}	0.25269	2.344	0.0229
Dw_{t-4}	0.32032	2.918	0.0052
Dp_{t-2}	0.38424	2.474	0.0166
Dsu_{t-1}	-0.17936	-2.711	0.0090
Dsu_{t-2}	0.16303	2.309	0.0248
Dsu_{t-3}	-0.19147	-2.772	0.0077
Dsu_{t-4}	0.06843	1.052	0.2976
$Seas_t$	0.03023	3.152	0.0027
Equation for Dp			
Dp_{t-1}	0.20314	2.210	0.0315
Dp_{t-3}	0.30286	3.375	0.0014
Dw_{t-3}	0.11756	2.345	0.0228
ecm_{t-1}	0.04339	2.166	0.0348
$Seas_t$	-0.02627	-4.575	0.0000
$Seas_{t-2}$	-0.05388	-9.976	0.0000
Constant	-0.10055	-1.637	0.1075
Equation for Dsu			
Dsu_{t-2}	0.27307	2.503	0.0154
Dsu_{t-4}	-0.26954	-2.660	0.0103
$Dhprod_{t-2}$	-0.62959	-3.401	0.0013
$Dhprod_{t-4}$	-0.52295	-2.768	0.0077
Dw_{t-1}	0.31497	1.826	0.0735
Dw_{t-4}	0.42025	2.402	0.0198
ecm_{t-1}	0.21837	3.138	0.0028
$Seas_t$	-0.05955	-2.473	0.0167
$Seas_{t-2}$	-0.06416	-2.765	0.0078
Constant	-0.65933	-3.027	0.0038

Table 10 (continued):

Equation for Dhprod			
$Dhprod_{t-1}$	-0.46372	-3.728	0.0005
$Dhprod_{t-2}$	-0.45514	-3.558	0.0008
$Dhprod_{t-3}$	-0.29770	-2.683	0.0097
Dw_{t-1}	0.19335	1.574	0.1215
Dw_{t-2}	0.24734	2.396	0.0201
Dw_{t-4}	0.28111	2.619	0.0115
$Seas_t$	-0.06179	-4.603	0.0000
$Seas_{t-1}$	-0.01465	-1.288	0.2034
$Seas_{t-2}$	-0.03981	-2.836	0.0065

history. The nominal wage inflation has also a positive effect which probably implies that, in the short-run, rises in the nominal wage inflation motivate rises in productivity growth.

The misspecification statistic results of the system are given in Table 11. The system can still be considered well specified even though there is evidence of increased serial correlation in the residuals for Dsu. Finally, in order to test if the chosen congruent simultaneous equations model parsimoniously encompasses the VAR, we performed a LR test for the overidentifying restrictions. The statistic which is asymptotically distributed as χ^2 (50), took the value of 44.287 (p-value: 0.7010) which provides evidence to accept the imposed restrictions.

Table 11: **SEM Diagnostic Statistics.**

Equation residual standard deviations

Dw	Dp	Dsu	Dhprod
0.02620	0.01119	0.03953	0.02593

Equation tests

Variable	Statistic	Value	p-value
Dw :	AR 1- 4F(4, 36) =	1.9455	[0.1240]
Dp :	AR 1- 4F(4, 36) =	2.5047	[0.0592]
Dsu :	AR 1- 4F(4, 36) =	3.8965	[0.0101]*
Dhprod :	AR 1- 4F(4, 36) =	2.5376	[0.0567]
Dw :	Normality $\chi^2(2)$=	2.4793	[0.2895]
Dp :	Normality $\chi^2(2)$=	3.7557	[0.1529]
Dsu :	Normality $\chi^2(2)$=	2.3434	[0.3098]
Dhprod :	Normality $\chi^2(2)$=	4.5972	[0.1004]
Dw :	ARCH 4 F(4, 32) =	0.1024	[0.9808]
Dp :	ARCH 4 F(4, 32) =	0.8933	[0.4793]
Dsu :	ARCH 4 F(4, 32) =	0.2499	[0.9076]
Dhprod :	ARCH 4 F(4, 32) =	0.2455	[0.9102]

Vector tests

VecAR	AR 1-4 F(64,135) =	0.7143	[0.9337]
VecN	$\chi^2(8)$=	12.214	[0.1419]

Table 11 (continued):

Correlation of residuals				
	Dw	Dp	Dsu	Dhprod
Dp	0.119	1		
Dsu	0.093	0.192	1	
Dhprod	0.238	0.098	0.068	1

6.2 Parameter constancy and forecasting

6.2.1 Parameter constancy

All four specifications (UVAR, PVAR, DVAR and SEM) obtain relatively constant parameters as evidenced by Chow tests (not shown for space reasons). However, they can also be compared according to the constancy of their parameters by making use of three forecast test statistics (see Doornik and Hendry (1994)).

The break point for the sample period is decided to be 1985.4. As shown by the graphs of the series at this period there was a shift to more restrictive economic policies which influenced seriously the behaviour of the series. In addition, the policy regime change had to be taken into account for the modelling of the system, by inclusion of a dummy for that period. Thus, the four alternative specifications were first estimated for the period until 1985.3 and then their dynamic forecasts over the period 1985.4 - 1990.2 were used for model comparison. The results of the one-step ahead forecast test statistics together with the means and standard deviations of the forecast errors are reported in Table 12. On the basis of these results the best overall performance is found in the SEM.

The SEM is the only model for which the parameter constancy assumption is not rejected by any of the obtained tests. Actual and forecast values for the PVAR, SEM and DVAR models are given respectively in

Table 12: **Testing for parameter constancy using forecast statistics**

UVAR: Period 1985 (4) to 1990 (2)	
F_1 using Ω	F(76,23)= 12.278 [0.0000]**
F_2 using V[e]	F(76,23)= 3.8368 [0.0003]**
F_3 using V[E]	F(76,23)= 2.4140 [0.0099]**
PVAR: Period 1985 (4) to 1990 (2)	
F_1 using Ω	F(76,21)= 2.9898 [0.0032]**
F_2 using V[e]	F(76,21)= 1.5351 [0.1345]
F_3 using V[E]	F(76,21)= 1.3046 [0.2506]
DVAR: Period 1985 (4) to 1990 (2)	
F_1 using Ω	F(76,22)= 2.6334 [0.0063]**
F_2 using V[e]	F(76,22)= 1.4129 [0.1819]
F_3 using V[E]	F(76,22)= 1.2903 [0.2549]
SEM: Period 1985 (4) to 1990 (2)	
F_1 using Ω	F(76,34)= 1.5700 [0.0731]
F_2 using V[e]	F(76,34)= 1.2738 [0.2192]

Table 12 (continued):

Descriptive statistics of forecast errors.

	Dw	Dp	Dsu	Dhprod
UVAR				
Mean	-0.0618	0.0138	-0.0591	0.0116
SD	0.0306	0.0243	0.0908	0.0280
PVAR				
Mean	-0.0407	-0.0003	0.0011	-0.0192
SD	0.0423	0.0181	0.0523	0.0353
DVAR				
Mean	-0.0293	-0.0028	-0.0103	-0.0106
SD	0.0408	0.0187	0.0567	0.0355
SEM				
Mean	-0.0075	0.0024	0.0162	0.0059
SD	0.0228	0.0132	0.0528	0.0301

Figures 7, 8 and 9.

6.2.2 Parameter constancy of the cointegrating relationship.

The SEM and PVAR models perform better than the DVAR one, in terms of parameter constancy. However, this may happen because of the way the models are specified. As Mizon (1995b) notices, this may be due to the fact that, even though possible regime shift are not taken into account explicitly by any of the three specifications (there is no dummy included in any of the models), the PVAR and SEM models include a full sample estimate of the ecm_t, which thus reflects the regime shift and keeps the forecasts on track. This would not happen, though, when comparing *ex ante* forecasts in case that a regime shift (which affects the long-run equilibrium mean) takes place in a time point after the analysed period.

Actually, as Clements and Hendry (1995), Hendry and Clements

Table 13: **Reduced sample weak exogeneity tests.**

1975.1 - 1985.3		
Hypothesis	$\chi^2(1)$	p-value
$\alpha_{11} = 0$: w. exogeneity for w:	3.047	0.0809
$\alpha_{21} = 0$: w. exogeneity for p:	10.61	0.0011**
$\alpha_{31} = 0$: w. exogeneity for $hprod$:	3.472	0.0624
$\alpha_{41} = 0$: w. exogeneity for su:	5.321	0.0211*

(1994) and Mizon (1995b) notice, the forecasts of the difference models remain unbiased when the long-run equilibrium mean has changed prior to forecasting due to an important regime shift. The models, though, which include the equilibrium correction terms (VECM, PVAR, SEM) will produce biased forecasts: the equilibrium correction terms tend to pull the forecasts towards the now inapropriate "equilibrium" [14].

It seems therefore necessary to perform cointegration analysis using the data sample before the break in 1985.4 and reestimate the PVAR and SEM models using the short sample long-run information in order to evaluate their *ex ante* forecasting performance.

Cointegration analysis performed for the period 1975.1 - 1985.3, gives evidence for two possible cointegrating relationships, one of which takes the form of a long-run relationship between real wage, unemployment and productivity:

$$w - 0.8455p + 0.0046su - 5.190hprod$$

The obtained cointegrating relationship is very close to the unrestricted cointegrating vector obtained by making use of the whole sample period, given in Table 5. The hypothesis of cointegration between real hourly wage, hourly productivity and unemployment is accepted for a LR $\chi^2(1)$

[14]Hendry argues in his co-breaking theory (1996), that a solution to this problem could be the exploration of whether and how the regime shifts that occur in a number of variables, are related.

test value of 2.8913 (p-value = 0.0891) and gives a relationship of the form:

$$w - p - hprod - 0.1673su$$

which can be used as an error correction term, *ecm*1. In the reduced sample cointegrating vector, it is just the magnitude of the coefficient of su that changes, with no change in the sign. In addition, tests for weak exogeneity of the variables with respect to the long-run parameters reveal no change in their status when the reduced sample is used. The results reported in Table 13 indicate that unemployment and prices remain the endogenous variables of the relationship.

In addition, the assumption that b takes the value -0.11 obtained by the whole sample analysis, is accepted when tested for the period 1975.1 - 1985.3: The relevant LR test distributed as a $\chi^2(2)$ takes the value = 4.516 (p-value= 0.1045).

In Figure 7 the graphs of the two cointegrating vectors obtained for the different periods can be compared visually. All evidence support that the policy change did not have a very strong effect in the long-run behaviour of the variables.

6.2.3 Forecasting comparison.

In a final step, the short sample cointegrating vector *ecm*1 replaces *ecm* in the PVAR, forming PVAR1 and in SEM specification forming SEM1, and the *ex ante* forecasts are compared with those of the DVAR. The forecast test results and the means and standard deviations of the forecast errors are reported in Table 14. The new model SEM1 again has the best forecasting performance among the three models DVAR, PVAR1 and SEM1. Actual and forecast values for SEM1 are given in Figure 11.

Table 14: **Forecasting tests.**

PVAR1 forecasting: Period 1985 (4) to 1990 (2)

F_1 using Ω	F(76,21)=	3.3188	[0.0015]**
F_2 using V[e]	F(76,21)=	1.6064	[0.1106]
F_3 using V[E]	F(76,21)=	1.3329	[0.2325]

SEM1 forecasting: Period 1985 (4) to 1990 (2)

F_1 using Ω	F(76,34)=	1.7018	[0.0437]*
F_2 using V[e]	F(76,34)=	1.3365	[0.1755]

Descriptive statistics of forecast errors.

	Dw	Dp	Dsu	Dhprod
PVAR1				
Mean	-0.04334	0.00146	0.00688	-0.02360
SD	0.04109	0.01809	0.05367	0.03438
SEM1				
Mean	-0.00744	0.00453	0.03149	0.00587
SD	0.02287	0.01314	0.05330	0.03002

7 Conclusions

Price and wage determination in Greece was investigated using labour market theories describing wage setting and the relationship between wage and price inflation. The sample period included different political regimes with different weights on inflation control, the effects of two devaluations of the national currency at 1983 and 1985 and the beginning of the EEC membership at 1981. The analysis was done in a closed system framework including the variables prices, hourly wages and hourly productivity in manufacturing and unemployment.

A thorough investigation of the time dependence properties of the series on a univariate level, indicated that the unemployment series contains a strong and unstable seasonal component and is seasonally integrated. These properties led to the use of a seasonally adjusted series of unemployment for the modelling of the system. The empirical analysis followed the "encompassing the VAR" methodology, according to which simultaneous equations models are evaluated by their congruence and their ability to encompass the VAR congruent representation of the data generation process. In addition, the Johansen cointegration analysis which takes into account the nonstationarities of the series on a multivariate level, and provides a framework for the joint analysis of long-run and short-run behaviour was used.

A long run positive real wage - productivity relation with positive unemployment level effects, in which price and unemployment are the endogenous variables, was established. The positive unemployment level impact probably reflects insiders - outsiders effects, real wage rigidities and inability of the productive sector to react to positive shocks. The result came out from the long-run analysis of the labour market, where alternative theoretical hypotheses, including stationarity for the individual series, were tested. Then, the long-run information was incorporated in a reduced, yet congruent parameterisation of the initial system, (the PVAR), which has been used as the benchmark within which alternative models were evaluated. The finally chosen simplified model, SEM is shown to be congruent and able to encompass the PVAR. It has been given reasonable theoretical interpretation and has constant parameters. In addition, it gives better forecasts than the DVAR model (which can be considered a powerful rival model within the time series analysis tradition) even when it is estimated by using the cointegrating relation obtained for the period before the policy change regime characterising the whole period, took place.

APPENDIX: Data definitions and sources.

- Y = Index of industrial production in manufacturing. Source: OECD Main Economic Indicators, various issues (OECD).
- E = Employment in manufacturing. Source: OECD.
- W = Nominal hourly earnings in manufacturing. Source: OECD.
- P = Consumer price index. Source: International Financial Statistics, International Monetary Fund (IMF).
- H = Weekly hours of work in manufacturing industry. Source: OECD.
- U = Number of registered unemployed. Source: OECD.

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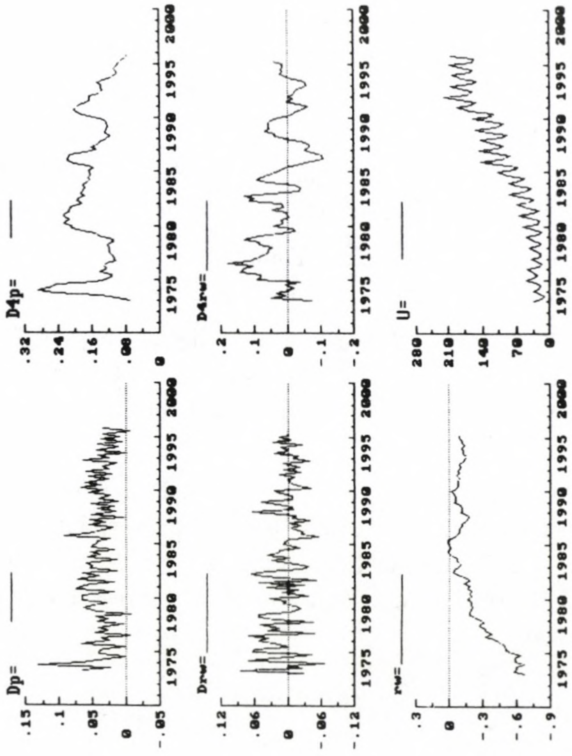

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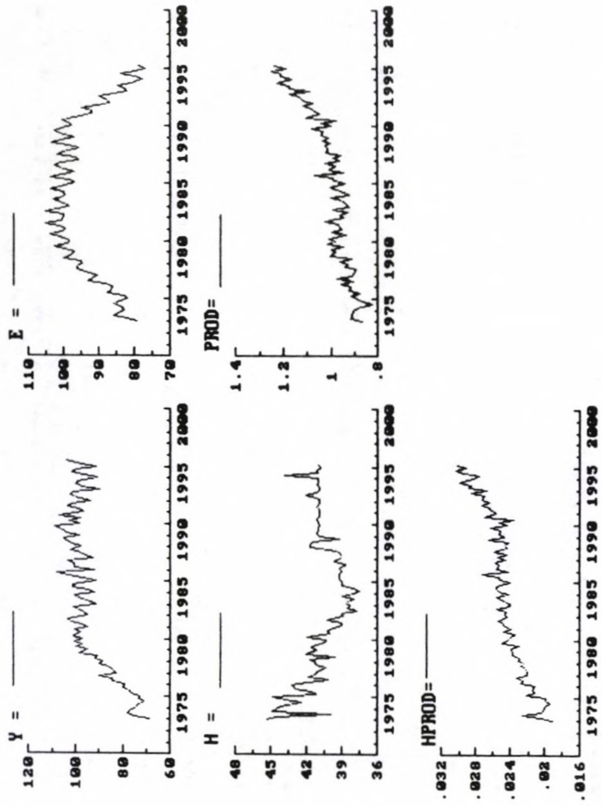

Figure 2: Manufacturing sector series

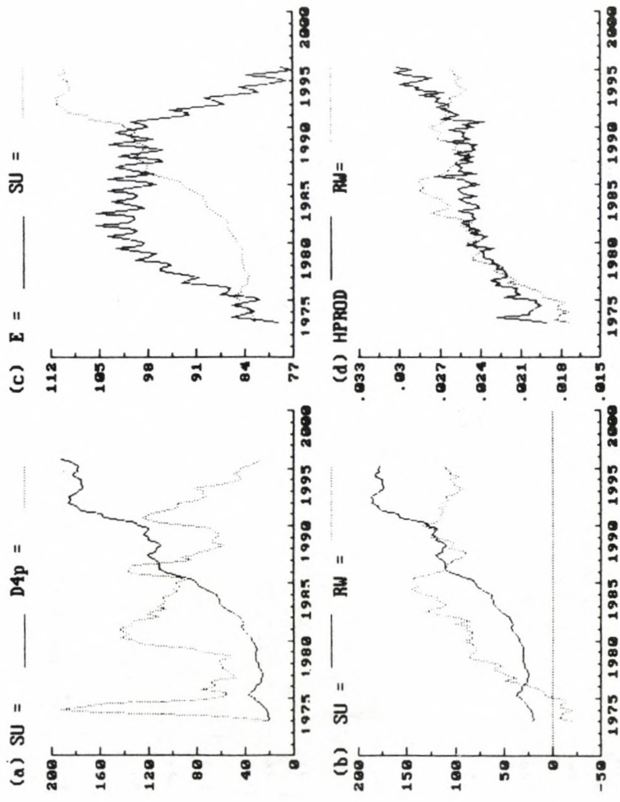

Figure 3: Mean and variance adjusted series

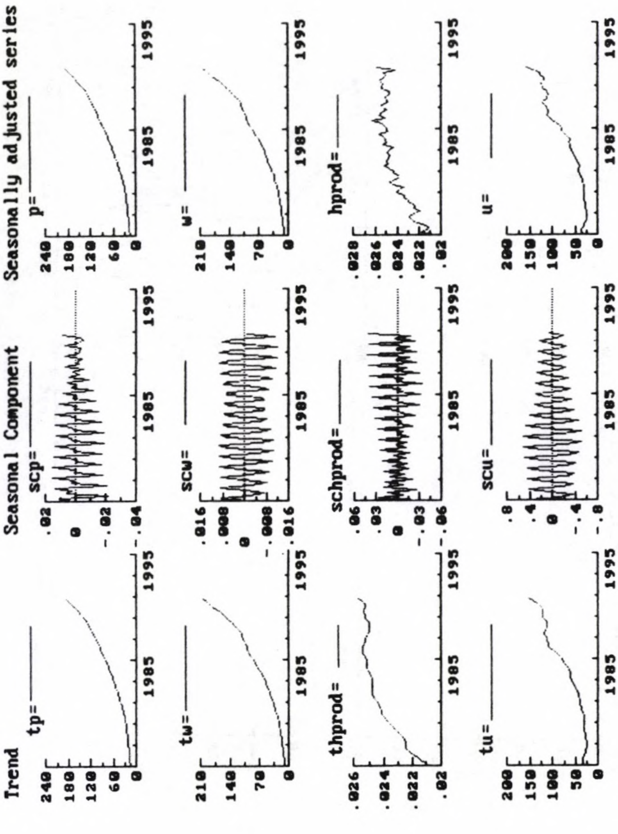

Figure 4: Trend, seasonal component and SA series

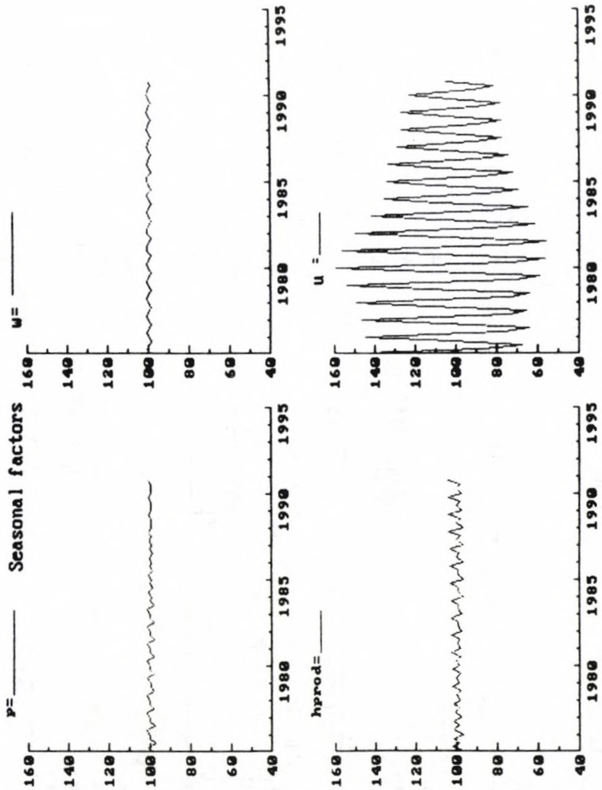

Figure 5: Seasonal factors of the series

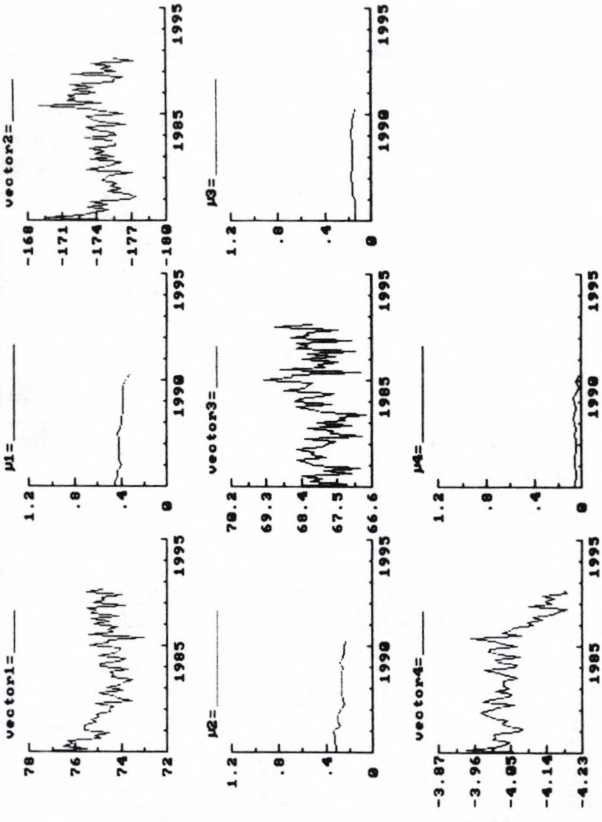

Figure 6: Cointegrating vectors and recursive eigenvalues

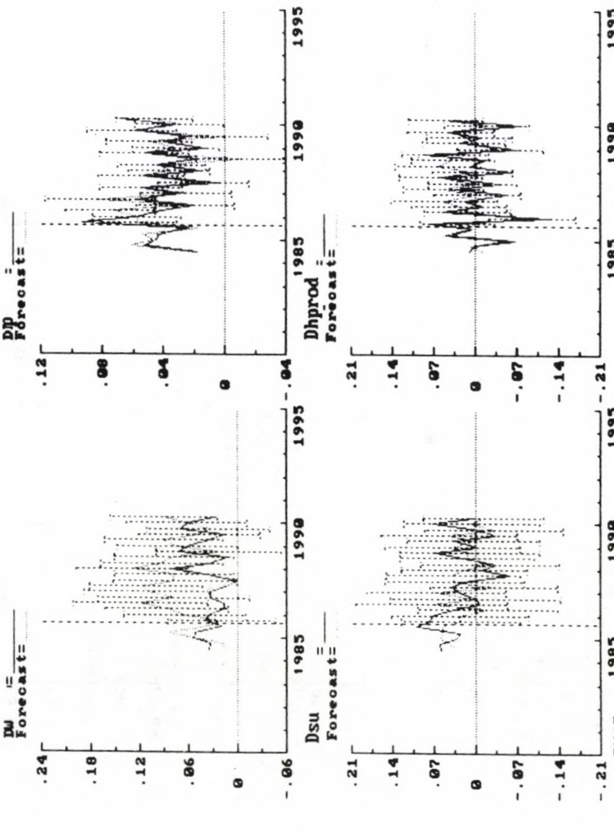

Figure 7: PVAR forecasts

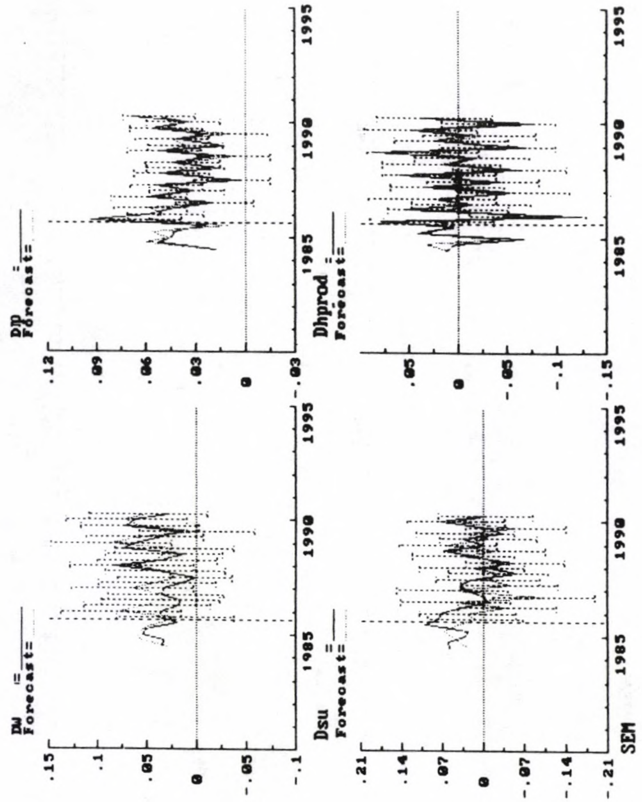

Figure 8: SEM forecasts

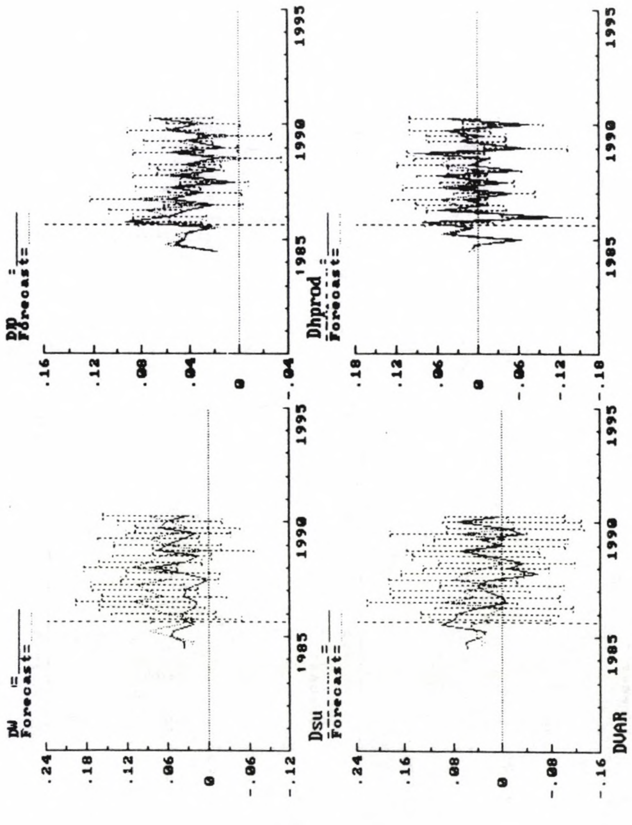

Figure 9: DVAR forecasts

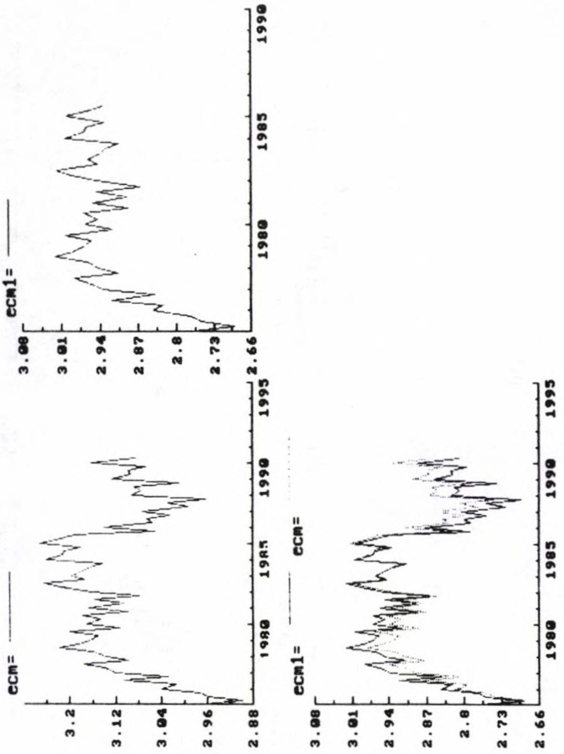

Figure 10: ecm, ecm1

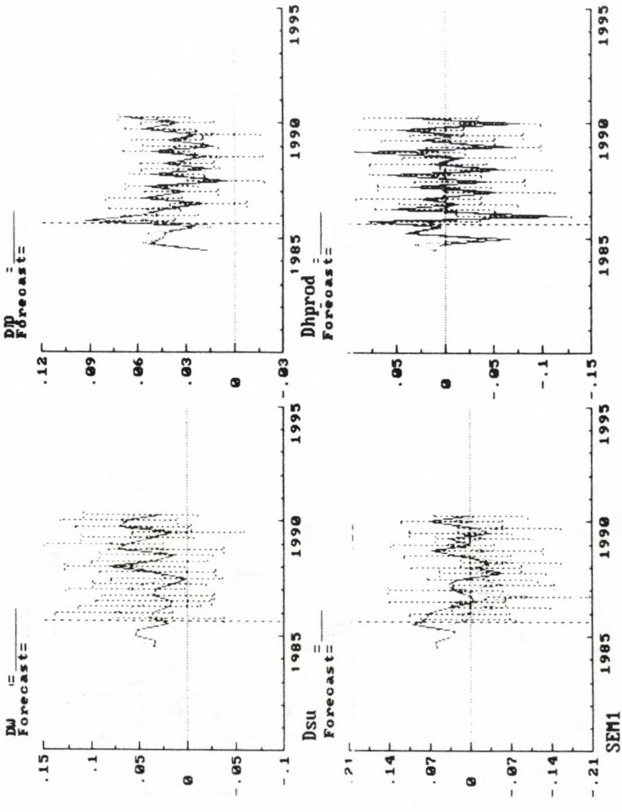

Figure 11: SEM1 forecasts

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